# Public Debt and Economic Growth:
# Panel Data Evidence for Asian Countries


## Abstract

This study examines the relationship between public debt on both short and long-run economic growth, in a panel of selected Asian countries for the period of 1980-2012. We employ several econometrics methods: pooled mean group, mean group, dynamic fixed effects and also allow for common correlated effects. The impact of a change in public debt is also analysed using asymmetric panel ARDL method. Our results indicate that an increase in government debt is negatively associated with economic growth in both the short and long-run.



**Keywords:** public debt, economic growth, dynamic panel data, asymmetries, ARDL

**JEL classification:** F33, F34, F35, O11



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## 1. Introduction

The relationship between public debt and economic growth has been a subject of increasing interest amongst academic scholars and policy makers. The problem of rising public debt is nothing new to developed countries and has also been an issue of increased interest in developing countries. Reinhart and Rogoff (2010) argue that a public debt to GDP ratio above 90% is associated with a lower economic growth rate. However, recent empirical studies such Égert (2015) question this result and conclude that no simple public debt threshold exists. Indeed, in their study Herndon *et al* (2014) conclude that there are significant errors in the results of Reinhart and Rogoff and that the 90% adverse debt threshold impact on economic growth is non-existent. Also, Reinhart and Rogoff do not deal with the issue of causality and Dafermos (2015) shows that their results are heavily impacted by periods of low economic growth in which there is usually a noticeable increase in public debt. There are two mechanisms through which this happens: (i) low economic growth directly impacts on the debt to GDP ratio since GDP is the denominator of this ratio and (ii) low economic growth tends to worsen fiscal deficits due to the impact of automatic stabilisers.

This study analyses whether rising public debt is harmful for growth, in both the short-run and long-run using data from fourteen Asian countries. The fact that the Asian countries are the biggest borrowers among emerging economies means that the issue of rising public debt is a particularly important issue in that region of the world. The Asian economies have also been exposed to two major crises during the period under study, the Asian financial crisis of 1998 and the global financial crisis of 2008 which has boosted the public debt to GDP ratios in these countries.

We firstly use a simple bivariate model to assess the direct impact of government debt on economic growth. We then use the panel autoregressive distributed lag (ARDL) of pooled mean group (PMG), mean group (MG) and dynamic fixed effects (DFE) to examine the relationship. The asymmetric panel ARDL is used to examine the different response of change in debt over the short-run and long-run relationship, this technique is also used in an analysis of debt in the Eurozone Area by Gómez-Puig and Sosvilla-Rivero (2015). This study contributes to the literature by examining a set of countries that has not previously been explored, using both a linear and non-linear methodology. Our research also contributes by looking at the issue of cross-sectional dependence in macro panel data. The presence of cross-sectional dependence may be caused by numerous aspects: spatial spillover, omitted and unobserved common factors as discussed in Breitung and Pesaran (2008). Ignoring these

factors may lead to the inconsistency of parametric and nonparametric estimators as pointed out by Baltagi (2014).

The remainder of this study is organised as follows: section 2 presents a review of the literature, section 3 explains the data and sample selection, section 4 presents the empirical methodology and section 5 discuss the empirical results. Finally, section 6 concludes.

## 2. Literature Review

Debt-related problems are nothing new to developed and developing countries. Many different empirical approaches have been used to examine the link between public debt and economic growth. The results can be heavily influenced by the time period of the study, country selection and estimation methods. As pointed out by Panizza and Presbitero (2013) at the theoretical level models yield ambiguous results concerning the relationship between public debt and economic growth and hence the link between the two is basically an empirical issue. They also argue that there is no empirical study that can make a compelling case for a causal relationship going from debt to economic growth. Nonetheless, a strong adverse effect of high levels of public debt in low-income countries is found in the study of Pattillo *et al.* (2004) who find that for highly indebted nations, per capita growth will be reduced by one per cent if debt is doubled. Cecchetti *et al.* (2011) based on an analysis of 18 OECD economies argues that there is a threshold at 85% debt to GDP which is exceeded leads to a reduction in future economic growth. They find that after the threshold that an increase of 10% in government debt to GDP ratio will lower annual economic growth by 0.17-0.18% over the following 5-year period.

Theoretically, government debt can stimulate a nation's long-run growth if it is productively allocated to the determinants of growth. Several studies argue that there is a non-linear link in which increased government debt may increase economic growth initially but after a certain point lead to a decrease in the growth rate. The frequently cited study of Reinhart and Rogoff (2010), finds a weak link between low public debt levels and growth but argues that if debt-to-GDP ratio exceeds 90%, it is harmful to growth. Based on historical data series for two decades, they analyse the link among inflation, high central sovereign debt, and economic growth in both developed and developing nations. However, the Rogoff and Reinhart results have become controversial due to some computational errors in their calculations that were pointed out by Herndon *et al.* (2014) that undermine the existence of 90% threshold in the debt-growth link, arguing no discontinuity above the cut-off level. Another replicative study by Minea and Parent (2012) shows that the threshold point is somewhat higher than 90%

of GDP and kicks in at around the 115% debt-to-GDP ratio above that ratio public debt is found to be negatively correlated with growth.

Despite the controversy, other research provides some support to Reinhart and Rogoff's (2010) findings. Kumar and Woo (2010) find a similar result using panel data for 38 advanced and emerging countries. They analyse the impact of initial government debt on subsequent per capita GDP growth and show that there exists an inverse-U shaped relationship with a more than 90% debt-to-GDP being detrimental to economic growth. Checherita-Westphal and Rother (2012) investigate the Euro area's economic growth using annual and cumulative five-year overlapping data of government debt and a squared debt term to analyse the non-linearity of the relationship and conclude that after around the 90-100% debt to GDP ratio there is an adverse effect of public debt on economic growth. However, the quadratic relationship is very sensitive to extreme values, particularly in a small sample of observations as pointed out by Panizza and Presbitero (2013). Presbitero (2012) uses total government debt in analysing the debt-growth link in developing countries for the period 1990-2007. He finds conditional convergence and a threshold effect for debt-to-GDP ratio above 90% which is similar to some of the studies of advanced nations.

Apart from the non-linearity debate, the issue of reverse causality needs to be taken into consideration, that is, whether debt leads to higher growth or *vice-versa*. Many studies are critical of Reinhart and Rogoff's (2010) as a comparison because their research failed to consider the endogeneity issue. Checherita-Westphal and Rother (2012) uses various instrumental variable (IV) methods with two-stage least squares or GMM estimators. Their results suggest that the two-step GMM is more favourable regarding efficiency. Kumar and Woo (2010) use a system GMM dynamic panel regression approach to address the endogeneity issue. Both studies use a dataset in which cross-section is larger than the time span (N>T). The GMM method is considered to be more efficient and give more precise estimations since this approach is applicable for large cross-country analysis (see Roodman, 2009). Panizza and Presbitero (2012) also consider the foreign currency debt as a proportion of government debt, as an instrumental variable. However, the use of this variable is questionable, in terms of the economic interpretation and according to Woo and Kumar (2015) this variable cannot meet the restriction criteria of a good IV estimator and its usage as an instrument is highly questionable for high-income countries where there is a low level of foreign currency portion of debt.

To avoid reverse causality, Woo and Kumar (2015) use initial debt levels to analyse the effect on future growth. Due to the problem of finding suitable external instrumental variables, the standard system GMM estimator is used to address the potential endogeneity issue. They

find that a high initial level of public debt is significantly associated with slower subsequent growth in a large panel of countries made up of developed and emerging market economies. While Baum *et al.* (2013) show a positive short-run impact of public debt on growth, but the impact becomes smaller once the debt-to-GDP ratio is above 67%. They use the dynamic panel method of GMM to estimate the linear model and modified Caner and Hansen (2004) approach to estimate the debt threshold. Sen *et al.* (2007) also exploit a dynamic GMM to study the behaviour of government debt on economic growth in Asian and Latin American countries. In the spirit of debt overhang, they examine external debt and find that borrowing severely hinders growth in Latin America and has a mildly negative effect in the case of Asia.

However, the GMM only captures the dynamics of short-run and ignores the long-run relationship since the estimator is designed for small time span. Consequently, as shown by Christopoulos and Tsionas (2004) the outcomes may show a spurious result instead of long-run equilibrium. Moreover, in the case of small N and large T, the GMM estimator may suffer from an autocorrelation problem in the residuals of the first-difference estimation, see Roodman (2009). Gomez-Puig and Sosvilla-Rivero (2015) use ARDL bound testing when estimating debt-growth links in the Economic and Monetary Union countries. Focusing on time series estimation, the authors find that the adverse impact is persistent in the long-run, but there are positive effects for some member countries in the short-run. Conversely, Eberhardt and Presbitero (2015) use a dynamic model of common correlated effects of pooled group and mean group estimators to analyse the link between debt and growth and they also use the traditional mean group and dynamic two-way fixed effects as a means of comparison. Using data from 118 countries, the authors allow for heterogeneity in the long-run and short-run link. They find a significant positive effect on average in the long-run debt but insignificant result in the short-run.

The use of panel autoregressive distributed lag (ARDL) models for analysing the impact of public debt on economic growth can also be found in Chudik *et al.* (2017) using data on a sample of 40 developed and developing countries over the period 1965-2010 they find an adverse effect of increases in the public debt to GDP ratio on economic growth. They also find no simple debt threshold for either developed or developing countries after accounting for the impact of global factors and spillover effects.

## 3. The Data Set

Panel estimation is chosen in this study to control for individual heterogeneity, to identify unobservable characteristics and to give more information on reliable estimation, see Baltagi,

(2005). Our analysis uses the data of 14 countries in Asia over a period of 33 years (1980-2012), resulting in a total of 462 observations (see Table 1 for countries in the sample). The choice of the countries was determined by issues of data availability. Japan was excluded from the analysis due its high public debt level. Since the data consists of a panel of 14 countries for 33 years, where N=14, is much less than T=33 the GMM estimator is not appropriate for our analysis. Table 1 provides comparative data for countries debt-to-GDP ratio. However, when T is larger than N (as in our case) the ARDL approach is appropriate and therefore is the preferred method for our analysis.

**Table 1: Comparative Features of Public Debt in Asian Countries**

| Categories | Countries | Public Debt-to-GDP (%) | |
|---|---|---|---|
| | | 1995 | 2010 |
| Lower-middle Income | Bangladesh | 51.3 | 38.5 |
| | India | 69.5 | 67.0 |
| | Indonesia | 32.0 | 26.8 |
| | Nepal | 63.4 | 35.4 |
| | Pakistan | 68.0 | 61.5 |
| | Philippines | 62.7 | 43.5 |
| | Sri Lanka | 92.0 | 81.9 |
| Upper-middle and High Income[1] | Thailand | 12.2 | 41.9 |
| | Turkey | 34.6 | 42.3 |
| | Iran, I.R. of | 35.2 | 16.7 |
| | China: Mainland | 6.1 | 33.5 |
| | Malaysia | 41.6 | 53.5 |
| | Republic of Korea | 7.1 | 33.4 |
| | Singapore | 70.1 | 101.7 |

Data is obtained from different sources of macroeconomic variables: the public debt-to-GDP ratio is derived from the study of 'A Historical Public Debt Database' by Abbas *et al.* (2010) in conjunction with the World Development Indicators (WDI-WB) of the World Bank (2016) and official national statistics. Other control determinants are taken from the WDI-WB and the Penn World Tables 9.0 (PWT 9.0) (Feenstra *et al.*, 2015). Following Sala-i-Martin *et al.* (2004), the explanatory variables are a set of determinants of economic growth. With the

[1] Republic of Korea and Singapore are classified as high-income countries since 1997 (Nielsen, 2011)

inclusion of several control variables to overcome the problem of omitted variables bias. The variables used in our study are listed below:

- Real GDP (in log) is obtained from the PWT 9.0.
- Public debt (in log) is obtained from Abba et al (2010), the WDI-WB and official national statistics
- Average years of schooling (in log) our proxy for human capital, following several public debt-growth literature (Pattillo *et al.*, 2004; Woo and Kumar, 2015) is obtained from the PWT 9.0.
- Trade openness (in log): This study uses sum of import and exports as a percentage of GDP to accounts for international trade activity. Data is obtained from the WB-WDI.
- Investment ratio (in log) is obtained from the WB-WDI using gross fixed capital formation as a percentage of GDP.

## 4. Methodology

We use several econometrics methods to examine the relationship between public debt and economic growth particularly in Asian countries and consider both the long-run and short-run relationships, along with the presence of nonlinearity. To examine the short and long-run relationships we use the panel ARDL initiated by Pesaran and Smith (1995) and Pesaran *et al.* (1999). To solve the contemporaneous correlation issue, CS-ARDL of Common Correlated Effect Pooled Mean Group (CCEPMG) and Common Correlated Effect Mean Group (CCEMG) are used. Since the debt levels are generally well below the 90% we do not look for threshold effects in our analysis.

### *4.1. Preliminary tests*

We first conduct panel unit root tests before performing the main estimations, the tests are necessary to check whether the variables are non-stationary. Several tests are conducted: Im *et al.* (2003) test (IPS), Levin *et al.* (2002) test (LLC) and second generation of IPS test (CIPS) of Pesaran (2007). The LL test is based on the assumption of non-heterogeneity of the autoregressive parameter, while the IPS test allows the heterogeneity while the CIPS unit root test relaxes the assumption of cross-sectional independence of the contemporaneous correlation All of these tests use the null hypothesis of non-stationarity. The selection of the lag length is chosen using the Bayesian-Schwarz criteria.

Another test we conduct is Cross-Sectional (CD) Pesaran (2015) which accounts for the presence of cross-sectional dependence. Panel data estimation assumes that disturbances are cross-sectionally independent, however, with the cross-country influences in the population, the issue of a cross-sectional link may arise. This dependence might be caused by similar geographical area, political or economic inducement (Gaibulloev *et al.*, 2014), therefore it is necessary to test the presence of cross-sectional dependence we also employ the CIPS and CD tests to check the residuals properties.

*4.2. Panel cointegration tests*

Two panel cointegration tests are employed here, based on the results of preliminary tests of non-stationarity. If the variables are non-stationary, then an examination for cointegration is conducted, using cointegration tests of Pedroni (1996) and Westerlund (2007). These cointegration tests are expected to reveal the existence or otherwise of a long-run relationship.

The Pedroni (1996) test proposes seven different panel cointegration tests to check the absence of cointegration. The seven-test relies on three between-dimension approaches and four within-dimension methods. Generalised least square correction is used to correct the independent idiosyncratic error terms across individuals. The Westerlund (2007) test exhibits four-panel cointegration estimation with the null of no cointegration, rejection of null hypothesis can be considered as the presence of cointegration in at least one individual unit.

*4.3. Dynamic panel ARDL tests*

Panel Autoregressive Distributed Lag (ARDL) is conducted if no-cointegration found from the previous methods. This method is superior regardless the underlying regressors exhibit I(0), I(1) or a mixture both (Pesaran and Shin, 1998) with a time span of over 20 years, the macro panel data method can be implemented. It was not appropriate to use the GMM estimator due to the nature of dataset. Following the extensive literature on dynamic panel data, we implement several estimators to assess the relationship between public debt and economic growth, by using Mean Group (MG), Pooled Mean Group (PMG), Dynamic Two-Way Fixed Effect (DFE) estimators (Pesaran and Smith, 1995; Pesaran et al., 1999).

The main model of panel ARDL approach is to obtain the relationship between public debt and economic growth:

$$y_{it} = \alpha_i + \sum_{l=1}^{p} \beta_0 y_{i,t-l} + \sum_{l=0}^{q} \beta_1 d_{i,t-l} + \sum_{l=0}^{q} \beta_2 x_{i,t-l} + u_{it} \quad (1)$$

By reparameterising equation (1):

$$\Delta y_{it} = \alpha_i + \Phi_i\left(y_{i,t-l} - \theta_1 d_{i,t-l} - \theta_2 x_{i,t-l}\right) + \sum_{l=1}^{p-1} \lambda_{il} \Delta y_{i,t-l} + \sum_{l=0}^{q-1} \lambda'_{il} \Delta d_{i,t-l} + \sum_{l=0}^{q-1} \lambda''_{il} \Delta x_{i,t-l} + u_{it} \quad (2)$$

with $i$ and $t$ represent country and time respectively, $y$ is the real GDP, $d$ is the public debt to GDP ratio, $x$ is a set of control variables: openness, human capital and investment ratio. Notation $\lambda, \lambda', \lambda''$ are the short-run coefficients of the lagged dependent variable, debt and other control variables respectively. The long-run coefficients are $\theta_1$ and $\theta_2$ for debt and other control variables. Lastly, $\Phi_i$ shows the speed of adjustment.

The PMG restricts long-run equilibrium to be homogenous across countries, while allowing heterogeneity for the short-run relationship. The short-run relationship focuses on the country specific heterogeneity, which might be caused by different response of stabilisation policies, external shocks or financial crises for each country. The MG estimator allows for heterogeneity in the short-run and long-run relationship. To be consistent, this estimator is appropriate for a large number of countries. For a small number of N, this method is sensitive to permutations of non-large model and outliers (Favara, 2003).

By contrast, the DFE estimator restricts the speed of adjustment, slope coefficient and short-run coefficient to exhibit non-heterogeneity across countries. Accepting this estimator as the main analysis tool requires the strong assumption that countries response is the same in the short-run and long-run, which is less compelling. Another drawback is that this approach may suffer from simultaneity bias in a small sample case due to the endogeneity between error term and lagged explained variable (Baltagi *et al.*, 2000).

In the case of our data it is derived from middle-income countries which exhibit similar behaviour in the long-run, regarding economic growth. The short-run is expected to be non-homogenous due to the country specific differences, as such the PMG estimator seems to be superior to other methods. We use the Hausman test to verify the significance of each estimator. One important point is that ARDL, especially PMG and MG estimators, can alleviate the problem of endogeneity with the inclusion of sufficient lags of all variables (Pesaran *et al.*, 1999).

The common correlated effect is introduced in the panel ARDL estimation to account for contemporaneous correlation. By creating indicators of (weighted) cross-sectional averages

of regressors to control the common factor, this study focuses on Common Correlated Effect Pooled Mean Group (CCEPMG) method and adds Common Correlated Effect Mean Group (CCEMG) as a comparison (Pesaran, 2006). It is expected that CCEPMG to be consistent and efficient in this estimation, under the null hypothesis of no heterogeneity in the long-run.

*4.4. Asymmetric panel ARDL tests*

The nonlinear ARDL estimator of Shin *et al* (2014) allows for an asymmetric short-run and long-run relationship. Following Eberhardt and Presbitero (2015), this study attempts to look at the asymmetric response of long-run and short-run response of public debt accumulation in economic growth.

*Asymmetric panel ARDL tests*

Asymmetric long-run estimation requires a decomposition of variable of interest into its positive and negative sub-variables, which define ($\boldsymbol{d}^{+}$) and ($\boldsymbol{d}^{-}$) as partial sums of positive and negative changes in public debt.

$$y_{it} = \alpha_i + \sum_{l=1}^{p} \beta_0 y_{i,t-l} + \sum_{l=0}^{q-1} \left(\beta_1 d_{i,t-l}^{+} + \beta_2 d_{i,t-l}^{-}\right) + \sum_{l=0}^{q} \beta_3 x_{i,t-l} + u_{it} \tag{3}$$

where $d_{i,t}^{-} = \sum_{j=1}^{t} \Delta d_{ij}^{-} = \sum_{j=1}^{t} max\left(\Delta d_{ij}, 0\right)$ and $d_{i,t}^{+} = \sum_{j=1}^{t} \Delta d_{ij}^{+} = \sum_{j=1}^{t} max\left(\Delta d_{ij}, 0\right)$

By reparameterising equation (3) we obtain:

$$\begin{aligned} \Delta y_{it} = {} & \alpha_i + \Phi_i\left(y_{i,t-l} - \theta_1 d_{i,t-l}^{+} - \theta_2 d_{i,t-l}^{-} - \theta_3 x_{i,t-l}\right) \\ & + \sum_{l=1}^{p-1} \lambda_1 \Delta y_{i,t-l} + \sum_{l=0}^{q-1} \left(\lambda_2 \Delta d_{i,t-l}^{+} + \lambda_3 \Delta d_{i,t-l}^{-}\right) + \sum_{l=0}^{q-1} \lambda_4 \Delta x_{i,t-l} + u_{it} \end{aligned} \tag{4}$$

This study uses the PMG and the CCEPMG approach to account for cross-sectional dependence.

## 5. Empirical Results

We start our empirical analysis by conducting panel unit root tests for all our variables. The IPS and LLC unit root test assume cross-sectional independence, while the CIPS accounts for cross-sectional dependence. The unit root tests which are summarised in Table 2, show that variables of interest have both non-stationary and stationary characteristics. Real GDP, openness and human capital are I(1) according to all unit root tests. Investment is stationary

according to IPS and LLC tests, but considered non-stationary based on CIPS test. Government debt is stationary according to LLC, but has non-stationary characteristics based on IPS and CIPS. Consequently, it is necessary to perform cointegration tests between real GDP and public debt to GDP to check for the possible existence of a long-run relationship.[2]

| Table 2: Panel Unit Root Tests | | | | | | |
|---|---|---|---|---|---|---|
| | IPS Test | | LLC Test | | CIPS Test | |
| | Constant | Constant & Trend | Constant | Constant & Trend | Constant | Constant & Trend |
| **Tests in logarithmic levels** | | | | | | |
| Real GDP | 6.68 | 1.37 | 1.81 | -0.32 | -1.527 | -1.823 |
| Govt debt/GDP | -2.44 | -1.05 | -2.45*** | -3.75*** | -1.560 | -2.201 |
| Human capital | 2.51 | 0.29 | -0.55 | -2.45*** | -2.030* | -2.199* |
| Trade Openness | 0.56 | 0.01 | -0.21 | -0.88 | -1.802 | -2.106 |
| Investment ratio | -3.11*** | -2.86*** | -3.68*** | -3.92*** | -1.513 | -2.370 |
| **Tests in first logarithmic differences** | | | | | | |
| Real GDP | -10.49*** | -9.20*** | -10.52*** | -9.85*** | -3.892*** | -4.242*** |
| Govt debt/GDP | -11.52*** | -10.92*** | -- | -- | -4.417*** | -4.625*** |
| Human capital | -1.696** | -3.193*** | -1.96*** | -- | -- | -2.610*** |
| Trade Openness | -15.63*** | -14.57*** | -16.68*** | -15.31*** | -4.812*** | -5.062*** |
| Investment ratio | -- | -- | -- | -- | -4.611*** | -4.761*** |

***, **, * denote the rejection of the null of a unit root for 90%, 95% and 99% respectively.
Critical values: Without trend: -2.26 (1%), -2.11 (5%), -2.03 (10%)
With trend: -2.81 (1%), -2.64 (5%), -2.56 (10%)

Two cointegration tests are conducted to analyse the long-run relationship between government debt and growth. Pedroni test results (see Table 3) show that the null hypothesis of no cointegration in a heterogeneous panel cannot be rejected. To accept the alternative hypothesis the panel variance has to possess a large statistical value and the latter six tests have to show large negative values (Pedroni, 1996). The same result is obtained from Westerlund (2007) test of no cointegration between variables, showing high probabilities of no rejection in the *p-values*. The rationale here is to test for the absence of cointegration by determining whether an Error Correction Model (ECM) exists for individual panel members or for the panel as a whole. Two different classes of tests can be used to evaluate the null hypothesis of no cointegration and the alternative hypothesis: group-mean tests (*G*) and panel tests (*P*). Westerlund (2007) developed four panel cointegration test statistics ($G_t$ ,$G_a$, $P_t$ and $P_a$) based

[2] We have also checked for the presence of cross-sectional dependence across variables. We did this by conducting the Pesaran's cross-sectional dependence test. For all variables the tests rejected the null hypothesis of weakly cross-sectional dependence. Results of these tests are not reported here for economy of space and are available from authors upon request.

on the Error Correction Model (ECM).[3] The results of all these additional cointegration tests are summarised in Table 4 and in all cases show no evidence of cointegration.

**Table 3: Pedroni Cointegration Test Results**

| Variables (in log) | Test Statistics | Panel (Within Dimension) (a) | Panel (Within Dimension) (b) | Group (Within Dimension) (a) | Group (Within Dimension) (b) |
|---|---|---|---|---|---|
| Real GDP, | *V* | 0.2023 | -0.3559 | | |
| government | *rho* | 0.6958 | 1.2990 | 1.9790 | 2.4390 |
| debt/GDP | *t* | 0.1663 | 0.4502 | 1.1080 | 1.2580 |
| | *adf* | -0.7834 | 0.4393 | -0.7412 | 0.7053 |

Panel cointegration test include intercept and trend
*V*: the variance ratio, *t*: Pedroni test, *adf:* augmented dickey fuller
(a) Excluding dummy of banking crisis
(b) Including dummy of banking crisis

**Table 4: Westerlund Cointegration Test Results**

| Variables (in log) | Group and Panel Statistics | Constant (a) Value | Constant (a) *p*-value | Constant (b) Value | Constant (b) *p*-value | Constant and Trend (a) Value | Constant and Trend (a) *p*- value | Constant and Trend (b) Value | Constant and Trend (b) *p*-value |
|---|---|---|---|---|---|---|---|---|---|
| Real GDP, | $G_t$ | -0.074 | 0.991 | -1.628 | 0.992 | 0.183 | 0.993 | -1.574 | 0.910 |
| government | $G_a$ | 0.106 | 0.881 | -5.005 | 0.883 | 0.457 | 0.885 | -3.986 | 0.887 |
| debt/GDP | $P_t$ | 0.017 | 0.994 | -5.700 | 0.994 | 0.494 | 0.995 | -5.831 | 0.997 |
| | $P_a$ | 0.007 | 0.966 | -4.419 | 0.967 | 0.175 | 0.966 | -3.627 | 0.963 |

Note: $G_t$, $G_a$, $P_t$ and $P_a$ are defined in footnote 3.

As previously stated, the panel ARDL method can be utilised to account for long-run and short-run relationships, even for the case of non-stationary variables but without cointegration. Three methods are used in this study: PMG, MG and DFE. Table 5, Panel A reports the estimates for all three methods and shows a significant result in the short-run that increased government debt adversely affects economic growth in the bivariate model. However, none of these tests are significant in the long-run. The ECM has a significant negative sign for the error correction term which implies that this model converges to a long-run relationship.

The next estimation presented in Table 5, Panel (B) uses all the determinants of growth and shows a similar result as in the bivariate case model. In the short-run, three estimators show significant negative results of public debt to economic growth. The investment ratio has a

[3] $G_t = \frac{1}{N}\sum_{i=1}^{N} \frac{\hat{\lambda}_i^k}{s.e.(\hat{\lambda}_i^k)}$, $G_a = \frac{1}{N}\sum_{i=1}^{N} \frac{T\hat{\lambda}_i^k}{\hat{\lambda}_i^k(1)}$; $\hat{\lambda}_i^k(1) = \frac{\hat{\omega}_{ui}}{\hat{\omega}_{yi}}$ where $\hat{\omega}_{ui}$ and $\hat{\omega}_{yi}$ are the usual Newey and West (1994) long-run variance estimators. $P_t = \frac{\hat{\lambda}_i^k}{s.e.(\hat{\lambda}_i^k)}, P_a = T\hat{\lambda}_i^k$.

significant positive effect on the economic growth. However, although human capital and trade openness have a positive sign as expected they are largely not significant, except in the case of the human capital proxy variable in the MG estimator. The negative long-run relationship of public debt in the estimation is significant only in the PMG method. The other two estimators show a negative but insignificant sign. Investment and openness in the long-run are signed as expected but not significant. The DFE approach exhibits a significant positive effect of human capital in the long-run.

The error correction terms are again negative and significant showing convergence in the long-run. Among all of the error correction results, the highest speed of adjustment of 18.6% (-0.186) is derived from MG from the second model implying a correction of 18.6% for the discrepancy of the estimation.

As stated before, we expect the PMG estimator to be the best approach. PMG allows the short-run to have differing responses across countries, while it restricts the long-run to exhibit non-heterogeneity. One advantage of using the PMG is that for a relatively small cross section of data (14 countries) the PMG is less sensitive to the existence of outliers (Pesaran *et al.*, 1999). In addition, the problem of serial autocorrelation can be corrected simultaneously. The benefit of using panel ARDL with sufficient lags is a reduction of the problem of endogeneity (Pesaran and Smith, 1999) which has been a concern in the recent debt-growth literature.

This chosen estimator is valid only if the assumption of the long-run restriction is not rejected. As can be seen from Table 5, Panel B, the homogeneity restriction is efficient and significant under such a hypothesis. Moreover, the Hausman test for the first and second model reveals a preference for PMG approach. The residuals show an I(0) integration suggesting the regressions are not spurious.

**Table 5: Panel ARDL Estimation Results**

| Variable (in log) | | Panel A | | | Panel B | | |
|---|---|---|---|---|---|---|---|
| | | **PMG (a)** | **MG (a)** | **DFE (a)** | **PMG (b)** | **MG (b)** | **DFE (b)** |
| Long-run | Government debt | 0.624 | 1.027 | 0.592 | -0.008 ** | -1.244 | 0.101 |
| | | (0.355) | (1.017) | (0.329) | (0.237) | (0.855) | (0.109) |
| | Investment ratio | | | | 1.049 | -0.719 | 0.429 |
| | | | | | (0.734) | (0.978) | (0.261) |
| | Human capital | | | | 0.697 | 0.945 | 2.142*** |
| | | | | | (1.999) | (2.294) | (0.546) |
| | Trade openness | | | | 1.025 | 0.740 | 0.477 |
| | | | | | (0.762) | (0.698) | (0.238) |
| Short-run | Government debt | -0.136 *** | -0.125*** | -0.079*** | -0.099*** | -0.086*** | -0.059 ** |
| | | (0.027) | (0.027) | (0.008) | (0.025) | (0.027) | (0.008) |
| | Investment ratio | | | | 0.079 *** | 0.059 ** | 0.114*** |
| | | | | | (0.024) | (0.024) | (0.017) |
| | Human capital | | | | 0.302 | -0.955 | 0.021 |
| | | | | | (0.471) | (0.614) | (0.240) |
| | Trade openness | | | | 0.014 | 0.002 | 0.012 |
| | | | | | (0.016) | (0.009) | (0.015) |
| Error Correction Term | | -0.010 *** | -0.003 | -0.012*** | -0.021 | -0.186*** | -0.036 ** |
| | | (0.003) | (0.008) | (0.004) | (0.018) | (0.044) | (0.010) |
| CD | | 2.12 (0.03) | 3.40 (0.00) | 4.68 (0.00) | 2.20 (0.02) | 3.25 (0.01) | 3.95 (0.00) |
| AR | | 8.22** | 9.94*** | 4.18* | 4.01 | 1.93 | 1.9 |
| Residual | | I(0) | I(0) | I(0) | I(0) | I(0) | I(0) |
| RMSE | | 0.029 | 0.028 | 0.031 | 0.026 | 0.024 | 0.029 |
| Adjusted $R^2$ | | 0.808 | 0.833 | 0.768 | 0.866 | 0.906 | 0.803 |
| Number of observations | | 448 | 448 | 448 | 448 | 448 | 448 |

The estimation includes the dummy of banking crisis and constant

Number in parentheses is the standard error ***, **, * shows 1%, 5% and 10% of significance level

Despite the significant result of the variable of interests, the ARDL method disregards contemporaneous correlation across countries, which is caused by unobserved factors. Ignoring these factors can lead to less consistent of parametric and non-parametric estimators (Baltagi, 2014). This is shown from the CD test (Pesaran) result which indicates a high value of cross-sectional dependence in the error term and clearly rejects the null of weakly cross-sectional dependence. The contemporaneous correlation is expected to diminish when the common correlated model is introduced.

Table 6 shows the CCEPMG and CCEMG estimators for bivariate and multivariate models. In the bivariate model, the CCEPMG estimator shows a significant result in the short-run and long-run and error correction term. In the multivariate model, both estimators show a significant negative debt relationship in the long-run while neither is significant in the short-run although the error correction terms remains negative and appears to be a much higher value. The control variable of the investment ratio is positively associated in the short-run and long-run in the CCEPMG result showing that it is a key determinant of economic growth.

By contrast, the human capital coefficient is not significant in this estimation and has a negative sign in the short-run. This result is somewhat surprising given the idea that human capital is an important driver of economic growth. One possible explanation along the lines of Van Leeuwen (2004) is that average years of schooling is an imperfect measure of human capital, he argues that this variable cannot capture the efficiency of education. Moreover, since this variable is not expressed in terms of a monetary unit, it is not comparable with the capital stock formation monetary unit measurement. CCEPMG and CCEMG estimators show that trade openness is significantly negative in the short-run when it might be the case of trade liberalisation undermines domestic production due to import competition, see Gries and Redlin (2012). However, trade openness is positively associated with economic growth in the long-run, as shown in CCEPMG estimator.

All four tests show negative and significant result for the error correction term, supporting the evidence of a long-run relationship. When a deviation from the long-run exists, the speed of adjustment to the long-run equilibrium is derived from the absolute value of the error correction term. In the bivariate model, deviations can be corrected for at a rate of 16.4% (-0.164) based on CCEPMG and 25.9% (-0.259) according to CCEMG. In the multivariate growth model, the speed of adjustment is much higher at 70.4% (-0.704) and 74.6% (-0.746) according to the CCEPMG and CCEMG estimators.

**Table 6: Panel ARDL Estimation - with Common Correlated Effect**

| Variable (in log) | | CCEPMG (a) | CCEMG (a) | CCEPMG (b) | CCEMG (b) |
|---|---|---|---|---|---|
| Long-run | Government debt | -0.140*** | -5.460 | -0.105*** | -0.206*** |
| | | (0.032) | (5.175) | (0.015) | (0.055) |
| | Investment ratio | | | 0.079*** | 0.035 |
| | | | | (0.024) | (0.081) |
| | Human capital | | | 0.303 | -0.789 |
| | | | | (0.201) | (0.477) |
| | Trade openness | | | 0.093*** | 0.096 |
| | | | | (0.028) | (0.064) |
| Short-run | Government debt | -0.081*** | -0.051*** | 0.003 | 0.002 |
| | | (0.018) | (0.019) | (0.014) | (0.025) |
| | Investment ratio | | | 0.036 ** | 0.021 |
| | | | | (0.018) | (0.029) |
| | Human capital | | | -1.094 | -0.826 |
| | | | | (0.780) | (1.124) |
| | Trade openness | | | -0.054*** | -0.059** |
| | | | | (0.010) | (0.024) |
| Error Correction Term | | -0.164 *** | -0.259*** | -0.704*** | -0.746*** |
| | | (0.029) | (0.038) | (0.060) | (0.103) |
| CD | | -2.63 (0.00) | -2.05 (0.04) | -0.99 (0.32) | -1.65 (0.09) |
| AR | | 8.55** | 12.25*** | 1.13 | 16.84*** |
| Residual | | I(0) | I(0) | I(0) | I(0) |
| RMSE | | 0.023 | 0.022 | 0.018 | 0.016 |
| Adjusted $R^2$ | | 0.408 | 0.512 | 0.700 | 0.828 |
| Number of observations | | 448 | 448 | 448 | 448 |

Note: the estimation includes the dummy for the global financial crisis and constant
Number in parentheses is the standard error
***, **, * shows 1%, 5% and 10% of significance levels

The residual tests are I(0) for all estimations, it is worth noting that the CCE estimator is valid even in the presence of serial correlation in the error term (Pesaran, 2006). However, except for CCEPMG, the results still suffer from cross-sectional dependence. In order to be a valid estimator, CCEMG should satisfy two requirements (i) the number of cross-section averages should be at least equal to the number of unobserved common factors and (ii) sufficient lags of cross section averages, see Chudik and Pesaran (2015). However, including more lags of averages variables is not desirable in our case because of the relatively small sample size. The CCEPMG estimator is chosen as the preferred approach because of the econometric theory behind this estimator and the significance of outcomes in both models.

Besides, the estimator is correctly specified without the problem of autocorrelation and cross-sectional dependence. In general, the results point to the detrimental consequences of increased public debt in economic growth.

The results for asymmetric panel ARDL are reported in Table 7. To find the appropriate lag length in the estimation the general to specific method was used and the Wald test is employed to examine if there is an asymmetric short and long-run response of government debt changes on economic growth. The PMG estimator cannot distinguish the asymmetric link of change of debt in the short-run, as coefficient of government debt (+) and government debt (-) exhibits a negative sign. The Wald test cannot reject the null hypothesis of a symmetric link in the short-run and in the long-run. While the sign shows a change in magnitude in the short-run it is statistically insignificant and there is also a clear rejection of the existence of an asymmetric relationship in the long-run. Using the CCEPMG, the Wald test of long-run symmetry cannot be rejected, indicating that change in government debt does not affect the long-run relationship and using the same convergence rate to define the long-run growth. While in the short-run there is a significant asymmetry in the CCEPMG estimator. The stationarity in the residuals indicates that the results of the Wald test are not spurious and the CCEPMG estimator controls for the cross-sectional dependence. In addition, the error correction term is negative. Although the CCEPMG approach rejects the null of Wooldridge test of autocorrelation, the CCE method is still superior to the presence of autocorrelation in error term (Drukker, 2003; Pesaran, 2006).

The investment ratio is positive and significant in short-run analysis but does not show a significant association with economic growth in the long-run. The human capital shows positive but insignificant result in the long-run and a mix of negative and positive result in the short-run. Again, the trade openness exhibits a negligible negative impact in the short-run but a much higher positive but not significant influence in the long-run.

The negative impact of public debt in the short-run implies lower debt accumulation will lead to a higher economic growth. While the positive 1 percent growth rate of government debt will lower economic growth by -0.012 to -0.125 percentage points. In the long-run, the magnitude of the two different regimes is somewhat higher in the region od -0.091 to –0.132 percentage points indicating that an increase in public debt will lead to significant adverse effect on economic growth.

**Table 7: Panel Asymmetric ARDL Estimation**

| Variable (in log) | | PMG | CCEPMG |
|---|---|---|---|
| Long-run | Government debt (+) | -0.091 | -0.132 *** |
| | | (0.287) | (0.030) |
| | Government debt (-) | 0.525 | -0.098 *** |
| | | (0.515) | (0.035) |
| | Investment ratio | 1.449 | 0.046 |
| | | (1.005) | (0.030) |
| | Human capital | 2.059 | 0.552 ** |
| | | (1.470) | (0.280) |
| | Trade openness | 1.206 | 0.137 *** |
| | | (0.932) | (0.035) |
| Short-run | Government debt (+) | -0.125 *** | -0.012 |
| | | (0.048) | (0.017) |
| | Government debt (-) | -0.056 | 0.048 * |
| | | (0.031) | (0.028) |
| | Investment ratio | 0.089 *** | 0.056 ** |
| | | (0.024) | (0.022) |
| | Human capital | 0.105 | -2.335 *** |
| | | (0.453) | (1.147) |
| | Trade openness | -0.001 | -0.077 *** |
| | | (0.011) | (0.013) |
| Error Correction Term | | -0.017 | -0.711 *** |
| | | (0.013) | (0.074) |
| CD | | 1.68 (0.09) | -0.74 (0.45) |
| AR | | 2.36 | 5.24 ** |
| Residual | | I(0) | I(0) |
| RMSE | | 0.029 | 0.017 |
| Adjusted $R^2$ | | 0.890 | 0.752 |
| $W_{LR}$ (*p-value*) | | 1.04 (0.309) | 0.42 (0.520) |
| $W_{SR}$ (*p-value*) | | 1.04 (0.307) | 2.83* (0.094) |
| Number of observations | | 448 | 448 |

The estimation includes the dummy of banking crisis and constant
Number in parentheses is the standard error, unless stated otherwise
***, **, * shows 1%, 5% and 10% of significance level

## 6. Conclusions

The issue of public debt and its impact on economic growth has been an important topic of debate amongst academics and policymakers. This research contributes in the public debt-growth study by focusing on a selection of Asian countries that typically have public debt to GDP ratios well below those in developed countries . In general, our results suggest that public

debt has a detrimental effect on economic growth suggesting that the idea that the negative effects of public debt kick in only at ratios of public debt to GDP of 90% or more may not apply to the Asian economies. Our main findings can be summarised as follows: (i) there is a negative effect of the public debt ratio on economic growth, both in the short-run and long-run, (ii) the negative relationship is more significant when we use common correlated factors to address the issue of cross-sectional dependence, (iii) an asymmetric response of a change in public debt is found to be significantly negative in the short-run. As such, rises in short-run public debt negatively affect economic growth in the short-run but falls in public debt do not have a correspondinly positive effect on economic growth in the short-run.

The failure of the initial cointegration tests of Pedroni (1996) and Westerlund (2007) to detect a the long-run relationship, led us to resort to the use of more advanced methods examine the relationship. Using the panel ARDL approach, increased public debt can be shown to negatively affect economic growth in both the short and long run. This result does not change when allowing for common correlated effects in the analysis. An asymmetric response of a change in debt is significant only in the short-run, that is, an increase in public debt will have a negative effect on growth in the short-run while a decrease in public debt will not have a correspondingly positive short-run impact on economic growth but it is likely to do so in the long-run.

Our negative results in a set of countries that have relatively low public debt to GDP ratios complement the results of Pattillo *et al* (2004) and Fall *et al* (2015) who show the existence of quite low debt thresholds in emerging countries. Our results may also have some interesting policy implications. Firstly, there is a need to examine why increases in the public debt to GDP ratio have a negative effect on GDP growth in these economies; Is it because the increase in public debt is used to finance projects of little worth to future economic growth ? Or because it crowds out productive private investment ? Or is it because the increase in public debt has benefitted a few elites at the expense on increasing the debt burden on the rest of the population? The answer may be that a mixture of all three elements come into play. Secondly, the countries concerned should consider putting in institutional improvements and control mechanisms that ensure that increases in public expenditure that increase public debt, explicitly consider the likely impact on future economic growth. This could mean that much needed infrastructure projects are given priority over projects with little economic value added, such as, increased military and defence expenditure. Finally, there could be a greater focus in these countries on public sector expenditure evaluation, for instance, increases on public sector bureaucracy may not be as useful in promoting economic growth as greater public sector

expenditure on improving health and education systems. The paybacks from these various types of government expenditure should be explicitly modelled so as to increase the probability of a creating a positive link between increasing public debt and economic growth.

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